

1 Introduction

According to the standard model of tax competition, increasing mobility of firms induces a race to the bottom in corporate taxes.¹ Recent theoretical work has fundamentally questioned the relevance of this scenario. In most ‘new economic geography’ models, the strength of geographical agglomeration forces increases as goods and factors become more mobile. As a result, somewhat paradoxically, the scope for attracting firms through fiscal inducements could in fact shrink as technological and administrative obstacles to firm mobility are reduced. The existence of agglomeration forces could thus allow governments to continue to tax corporate income even once capital has in principle become highly mobile.

We provide an empirical assessment of the hypothesis that agglomeration forces can offset differences in corporate taxes as a determinant of firm location. Estimating location choice models for firm start-ups across Swiss municipalities, we find that high corporate taxes are indeed a deterrent to firm location, but that this deterrent effect is weaker for sectors that are more spatially clustered. Hence, agglomeration economies - be they due to externalities or to spatially concentrated endowments - can reduce the ability (and incentive) of jurisdictions to compete for firms via strategically low tax rates.

These results are based on Poisson regressions derived from firm-level profit functions in a location choice model. We first estimate a ‘baseline model’ of firms’ location choices, in which we introduce an explicit interaction term between municipal corporate taxes and a measure of sector-level agglomeration. In an alternative approach, we then estimate a ‘specific model’ that is formally derived from a model of spatial demand and supply conditions. In that model, the relative effect of taxes versus agglomeration forces features implicitly rather than via an explicit interaction term. We minimize simultaneity problems between taxes and firm location by using sector-level counts of *new* firms (or of employment by new firms) as the dependent variable, and municipal corporate taxes which apply identically to firms across all sectors as the independent variable. Unobserved sector characteristics are controlled for via fixed effects, and the qualitative results are shown to be robust across a range of specifications and at different levels of sectoral aggregation.

We proceed as follows. Section 2 provides a brief review of the relevant literature. Section 3 presents the two estimable models. Our empirical setting and data set are described in

¹For a comprehensive review of this literature, see Wilson (1999).

Section 4. Estimation results are reported in Section 5, and Section 6 concludes.

2 Literature background

The implications of agglomeration economies for strategic tax setting among jurisdictions competing for mobile tax bases have been studied in a number of theoretical contributions, including Ludema and Wooton (2000), Kind, Midelfart-Knarvik and Schjelderup (2000), Andersson and Forslid (2003), Baldwin and Krugman (2004), and Borck and Pflüger (2006).² The key insight of this literature is that agglomeration forces make the world ‘lumpy’: when capital (or any other relevant production factor) is mobile and trade costs are sufficiently low, agglomeration forces lead to spatial concentrations of firms that cannot be dislodged by tax differentials.³ In fact, agglomeration externalities create rents that can in principle be taxed by the jurisdiction that hosts the agglomeration.

New economic geography models can also accommodate configurations where agglomeration economies in fact add to the sensitivity of firm location to tax differentials because one firm’s location choice can trigger further inflows and thus the formation of a new cluster. In such knife-edge configurations, agglomeration economies exacerbate the intensity of tax competition (Baldwin *et al.*, 2003, Result 15.8; Konrad and Kovenock, 2008). We abstract from these theoretically conceivable but practically rather less likely situations to focus on configurations featuring established agglomerations. This is where the new economic geography implies qualitatively novel predictions for tax policy. The models typically feature a single increasing-returns sector, the intensity of whose agglomeration forces varies (mostly non-monotonically) with trade costs. Where agglomeration forces are strongest (i.e. at intermediate trade costs), the probability that the increasing-returns sector completely agglomerates in one region is highest, and the sensitivity to tax differentials is smallest.⁴ The greater is a sector’s observed spatial concentration, the larger, on average, are the underlying agglomeration economies, and the lower should be the sector’s locational sensitivity to tax

²See Baldwin, Forslid, Martin, Ottaviano and Robert-Nicoud (2003, ch. 15, 16) for a comprehensive overview.

³Conversely, Ottaviano and van Ypersele (2005) have shown that in the presence of agglomeration economies, tax competition can be second-best welfare-enhancing, as it may mitigate a tendency towards excessive spatial concentration of firms.

⁴More precisely, in the standard ‘core-periphery’ model the range of tax differentials that will not dislodge a given spatial allocation of firms is largest where agglomeration forces are strongest.

differentials.⁵

Empirically, there is considerable evidence to show that firm location is sensitive to differences in corporate taxes, across a range of methodological approaches.⁶ Moreover, Greenstone and Moretti (2004) have shown that it can ‘pay’ to attract new plants with fiscal inducements - both in terms of higher wages and higher property values. Greenstone, Hornbeck and Moretti (2008) found that such additional plants tend to raise the total factor positivity of incumbent neighboring plants, which in turn suggests that agglomeration economies are important.

Starting with Carlton (1983), numerous studies have estimated conditional logit models of firms’ location choices.⁷ The conditional logit model has the advantage of being formally derived from a firm’s stochastic profit function, but it becomes computationally demanding for estimation of models with individual data and large numbers of locational alternatives. Papke (1991) therefore suggested that location choice could also be represented by a region-level count model, such that estimation is based on maximum likelihood with an assumed Poisson distribution.⁸ The Poisson model was shown by Guimaraes, Figueiredo and Woodward (2003) to be formally equivalent to the conditional logit model with grouped data and group-specific fixed effects. We can therefore estimate the conditional logit model via Poisson, taking sectors as the grouping variable.

Devereux, Griffith and Simpson (2007) have previously explored the impact of agglomeration economies on the sensitivity to local fiscal incentives of firms’ location choices. They estimate a conditional logit model of plant location in Great Britain, including an interaction term of region-level fiscal incentives with the stock of pre-existing same-sector plants in the relevant region, and they find that fiscal incentives have a greater impact on attracting plants in regions with large stocks of existing plants. As fiscal incentives in British regions are negotiated individually for each proposed new establishment, unobserved plant-level features might

⁵Burbidge and Cuff (2005) and Fernandez (2005) have studied tax competition in models featuring increasing returns to scale that are external to firms, with firms operating under perfect competition. In these models, individual firm mobility is not constrained by agglomeration economies, and governments may compete even more vigorously to attract firms than in the standard tax competition model. Krogstrup (2008) shows that for tax competition to be intensified, external agglomeration economies must be relatively weak, in the sense that they are outweighed by dispersion forces that stabilize the overall spatial allocation of activity. Our working hypothesis is that agglomeration economies are sufficiently internalized by firms to affect firms’ locational sensitivity to tax differentials.

⁶See, e.g., Hines (1999) for a survey, and de Mooij and Ederveen (2003) for a meta-analysis.

⁷Recent applications include Guimaraes, Figueiredo and Woodward (2000), Figueiredo, Guimaraes and Woodward (2002), Crozet, Mayer and Mucchielli (2004), Head and Mayer (2004), Strauss-Kahn and Vives (2005), and Devereux, Griffith and Simpson (2007).

⁸Count models of firm location have subsequently been estimated by List (2001), Guimaraes, Figueiredo and Woodward (2004), and Holl (2004).

affect both the probability of a plant receiving a grant in a particular area and the probability of it locating in that area. Since statutory corporate taxes of Swiss regions are neither firm- nor sector-specific, our empirical setting does not present the estimation challenge affecting an analysis of the same question based on British regional grants.

Our analysis is novel in two additional ways, both motivated by a quest to tie our estimations closely to the theory. First, in our baseline specification, we focus on the interaction of taxes with a *sector*-specific measure of agglomeration, in order to capture the essence of the new economic geography insight on tax competition.⁹ As long as taxes vary within the bounds beyond which they would trigger discrete (‘catastrophic’) relocations of mobile sectors, the theory consistently suggests that stronger sector-level agglomeration forces imply a lower sensitivity of firm location to tax differentials.¹⁰

Second, in addition to a somewhat *ad hoc* baseline model similar to those that have typically been estimated in this literature, we construct a ‘specific model’ as a representative firm’s profit function that is formally derived from explicitly modeled spatial demand and supply conditions. In this model, the interaction between municipal taxes and sector agglomeration economies is not introduced by assumption but implied by considering agglomeration economies in the production function. The baseline model thus reflects a more statistical approach to the data, while the specific model incorporates more structure from economic theory. We shall show that both models lead to the same qualitative conclusions.

⁹Devereux *et al.* (2007), using a *location*-specific measure, show that it may be cheaper to attract a new plant to an existing cluster than to a peripheral location. This is an important and evidently policy-relevant result, but not what the theory necessarily predicts when the economy is in spatial equilibrium. In an interior spatial equilibrium with no relocation costs, expected profits at the locus of agglomeration (the ‘central’ location) and at the periphery are equalized. Whether a given change in fiscal inducements is then more effective at attracting firms to a central or to a peripheral location is indeterminate, as it depends on the functional form of the relationship between real returns and industry shares across locations. In the simulations reported by Borck and Pflüger (2006, Fig. 5), a given fiscal inducement will attract a larger number of firms if offered at the peripheral location than if offered at the central location.

¹⁰In ‘core-periphery’ models, which, in the absence of taxes, accommodate only perfectly agglomerated or perfectly dispersed spatial allocations of the mobile sector, marginal variations in relative tax burdens imply marginal reallocations of that sector among locations in the dispersed equilibrium but have no effect on sectoral location in the agglomerated equilibrium (see, e.g., Baldwin *et al.*, 2003). In models that accommodate partially agglomerated configurations even in the absence of taxes, more strongly agglomerated equilibria imply lower delocation elasticities (Borck and Pflüger, 2006).

3 Models of location choice

3.1 A baseline model: footloose and latent startups

At the most general level, there are two approaches to modeling the location of new firms. One approach is to consider an investor who has resolved to set up a firm somewhere among a given set of locations and then decides which location to pick. We refer to this as the ‘footloose startup’ model.¹¹ The other approach is to assume that potential entrepreneurs are spatially immobile and continuously decide whether or not to set up a firm.¹² We refer to this approach as the ‘latent startup’ model. To the empirical researcher, these two approaches are equivalent in two essential respects: the decision to set up a firm at a particular location is based in both cases on expected profits, and in both cases expected profits are best modeled as a combination of deterministic components and a stochastic term.

We posit a general profit function for a *footloose-startup* decision problem, where a firm belonging to sector i has decided to set up a new plant f and now considers which location j to choose (time subscripts implied):

$$\pi_{fij} = U_{ij}^B + \varepsilon_{fij} = \alpha_1 T_j + \alpha_2 A_i + \alpha_3 T_j A_i + \boldsymbol{\beta}' \mathbf{x}_{ij} + \varepsilon_{fij}. \quad (1)$$

U_{ij}^B summarizes the deterministic part of the baseline model (B) that is common to firms of a particular sector and at a particular location; T_j represents the relevant corporate tax burden at location j ; A_i represents the strength of agglomeration economies in sector i ; \mathbf{x}_{ij} is a vector of other variables that determine a firm’s profits in sector i at location j (such as factor prices, proximity to markets, etc.); α_1 , α_2 , α_3 and $\boldsymbol{\beta}$ are coefficients to be estimated; and ε_{fij} is a stochastic error term. In line with our data, the deterministic part U_{ij}^B does not include any firm specific information beyond the sector assignment. A sector’s propensity to agglomerate, A_i , may be determined by pecuniary and/or technological spillovers, or it may be due to the spatial concentration of immobile resources that are important to the sector.

Our interest is in the parameter α_3 : while we expect the attractiveness of a location j to fall in the level of its corporate tax burden, implying that α_1 should be negative, this sensitivity should be weaker in sectors that are subject to strong agglomeration forces. A positive α_3

¹¹This approach underlies the empirical literature on location choices using the conditional logit model.

¹²See, e.g., Becker and Henderson (2000) and Figueiredo *et al.* (2002).

would therefore confirm the result of the economic geography literature that agglomeration forces can offset industries' sensitivity to tax differentials.

If we treat the location decision problem as one of random profit maximization, firm f will pick location m if $\pi_{fim} > \pi_{fij} \forall j, j \neq m$. As shown by McFadden (1974), the assumption that ε_{fij} has an extreme-value type 1 distribution yields a simple expression for the probability of choosing location m : $p_{fim} = \exp U_{fim} * (\sum_j \exp U_{fij})^{-1}$. If we define a dummy variable d_{fij} that equals one if firm f chooses location j and zero otherwise, the log-likelihood of the conditional logit model becomes: $\ln L_{CL} = \sum_f \sum_j d_{fij} \ln p_{ij} = \sum_i \sum_j n_{ij} \ln p_{ij}$, where n_{ij} represents the number of firms in sector i that choose location j .

Guimaraes *et al.* (2003) have shown that the same log-likelihood, up to a constant, obtains if one assumes n_{ij} to be independently Poisson distributed. Thus, estimates obtained from a Poisson count regression of n_{ij} on all region specific and region-sector specific regressors plus a set of sector fixed effects are identical to those obtained from conditional logit estimation. We can therefore rewrite the random profit model (1) equivalently as follows:

$$E(n_{ij}) = \lambda_{ij} = \exp(\alpha_1 T_j + \alpha_3 T_j A_i + \beta' \mathbf{x}_{ij} + \gamma' \mathbf{d}_i), \quad (2)$$

where n_{ij} follows a Poisson distribution and \mathbf{d}_i is a set of sector dummies. The inclusion of sector dummies forces the control matrix \mathbf{x}_{ij} to consist exclusively of variables that vary across locations. The main effect of A_i , α_2 , is absorbed into the sector fixed effects.

The *latent-startup* model assumes that every location hosts a certain number of immobile actual and potential new firms ('entrepreneurs') per sector. At every point in time, each potential entrepreneur computes the net present value (NPV) from becoming active and uses this to decide whether or not to start an actual firm. This yields, for every location-sector pair, a supply and demand curve for new firms in birth-NPV space. The supply curve, which rises in NPV, depends primarily on the size of a location's pool of potential entrants. The demand curve traces how the NPV per firm changes as more firms become active in the same sector and location, and its position depends on variables such as local factor costs and local product demand. Total births are then determined by the intersection of these demand and supply schedules. Becker and Henderson (2000) show that, conditional on standard regularity conditions, this model leads directly to the Poisson specification (2). By employing Poisson estimation, we can therefore accommodate both the footloose and the latent startup models -

a considerable advantage given that it would be impossible based on available information to judge which of the two models represents a better approximation of the actual data-generating process.

3.2 A specific model for footloose startups

While the model of expected profits in equation (1) may be intuitive and general, it is not rooted in a formal representation of the firm's optimization problem. We now derive a profit function formally, drawing on a simple model proposed by Crozet, Mayer and Mucchielli (2004). This will lead to a particular specification of the profit function that can be viewed as an alternative to equation (1), thus offering a complementary framework for the exploration of our basic research question.

We assume identical consumer preferences but varying incomes across locations j , and we allow for different price elasticities of demand across sectors i . A generalized Cobb-Douglas utility function then implies the following expression for quantity demanded Q_{ij} :

$$Q_{ij} = \frac{\phi_i m_j^{\gamma_i}}{p_{ij}^{\delta_i}},$$

where ϕ_i is the sectoral expenditure share, m_j is relevant income at location j , p_{ij} is the price, γ_i is the income elasticity, and $\delta_i > 1$ is the price elasticity of demand.

Symmetry among firms of any sector at a particular location implies that quantity demanded, and thus equilibrium output per firm, are equalized: $q_{ij} = \frac{Q_{ij}}{N_{ij}}$, where N_{ij} is the number of active firms in sector i at location j .

Firms are assumed to be price takers in factor markets. Their unit costs are modeled as follows:

$$c_{ij} = (w_{ij} (1 + t_j^w))^{\theta_i^w} \left(k (1 + t_j^k) \right)^{\theta_i^k} (r_j (1 + t_j^r))^{\theta_i^r} N_{ij}^{(\theta_i^N + A_i)}, \quad (3)$$

where w_{ij} is the wage rate (which may vary across locations and industries), k is the capital rental price (assumed constant across locations and industries), r_j is land rental price (which may vary across locations), t_j^w is the payroll tax rate (to the extent that it is borne by employers), t_j^k is the capital tax rate, t_j^r is the property tax rate, A_i again captures agglomeration economies, and the θ s are parameters. θ_i^w , θ_i^k , and θ_i^r represent input shares of labor, capital

and land. The exponent on N_{ij} , $(\theta^N + A_i)$, implies that firms in more agglomerated sectors will benefit more from proximity to own-sector firms than firms in less agglomerated sectors.

Profits of a representative firm can be written as $\pi_{ij} = (1 - t_j^\pi)(p_{ij} - c_{ij})q_{ij}$, where t_j^π is the corporate income (i.e. profit) tax rate. Profit maximization with a large number of firms competing in quantities, and consideration of a multiplicative stochastic term ν_{fij} , implies the following firm-level profit function:¹³

$$\pi_{fij} = (1 - t_j^\pi) \phi_i m_j^{\gamma_i} \left((w_{ij}(1 + t_j^w))^{\theta_i^w} \right)^{1-\delta_i} \left((k(1 + t_j^k))^{\theta_i^k} \right)^{1-\delta_i} \left((r_j(1 + t_j^r))^{\theta_i^r} \right)^{1-\delta_i} N_{ij}^{((\theta^N + A_i)(\delta_i - 1) - 2)} \nu_{fij}.$$

In logs, this becomes:

$$\begin{aligned} \ln(\pi_{fij}) &= \ln(1 - t_j^\pi) + \ln \phi_i + \gamma_i \ln m_j \\ &+ ((1 - \delta_i)\theta_i^w) \ln w_{ij} + ((1 - \delta_i)\theta_i^w) \ln(1 + t_j^w) \\ &+ ((1 - \delta_i)\theta_i^k) \ln k + ((1 - \delta_i)\theta_i^k) \ln(1 + t_j^k) \\ &+ ((1 - \delta_i)\theta_i^r) \ln r_j + ((1 - \delta_i)\theta_i^r) \ln(1 + t_j^r) \\ &+ ((\theta^N + A_i)(\delta_i - 1) - 2) \ln N_{ij} \\ &+ \ln \nu_{fij}. \end{aligned} \tag{4}$$

We can thus write the following estimable equation:

$$\ln(\pi_{fij}) = U_{ij}^S + \ln \nu_{fij} \tag{5}$$

$$\begin{aligned} &= \beta_0 + \beta_{1i} + \beta_2 \ln(1 - t_j^\pi) + \beta_{3i} \ln w_{ij} + \beta_{4i} \ln r_j + \beta_{5i} \ln(1 + t_j^k) \\ &+ \beta_{6i} \ln(1 + t_j^w) + \beta_{7i} \ln N_{ij} + \beta_{8i} (A_i * \ln N_{ij}) + \beta_{9i} \ln m_j + \ln \nu_{fij}, \end{aligned} \tag{6}$$

where U_{ij}^S summarizes the deterministic part of the specific model (S), the ϕ_i are absorbed by sector fixed effects (β_{1i}), and property taxes t_j^r are dropped as they do not play a role

¹³See Crozet *et al.* (2004) for the derivation.

in our empirical setting.¹⁴ If we assume that $\ln \nu_{fij}$ follows an i.i.d. extreme-value type 1 distribution, equation (5) leads to a standard conditional logit model and can be estimated, *mutatis mutandis*, via a Poisson count model analogous to (2).¹⁵

The principal difference between the baseline model (2) and the specific model (5) is that the latter no longer features an explicit interaction term between the corporate income tax burden and sectoral agglomeration intensity. This stems from the simple fact that a given statutory tax rate on profits reduces profits exactly proportionally irrespective of any sectoral or locational specificities. As we will show in Section 5.3, the functional form of the conditional-logit estimate of equation (5) nonetheless implies that the marginal effect of the corporate tax varies with A_i .

3.3 Estimation issues

3.3.1 Scaling locations and firms

In estimating our models (2) and (5), we need to take account of the fact that real-world locations come in different geographic sizes, giving them different probabilities of attracting or generating a certain number of new firms even once their purely economic characteristics are accounted for. Failure to control for size differences could thus lead to omitted-variable bias.

We can think of our locations j as consisting of L_j equally sized ‘lots’ $\ell \in L_j$ each. Characteristics of individual lots are only observed (or are only relevant) at the level of locations.

Assume a conditional logit model *at the level of lots*. The profit function for firm f associated with lot ℓ in location j (suppressing industry-level notation) is given by $\pi_{f\ell} = U_{fj} + \varepsilon_{f\ell}$, where $\varepsilon_{f\ell}$ is independent across f and ℓ and follows an extreme-value type 1

¹⁴The profit function (4) implies that $\beta_2 = 1$. This restriction, however, cannot be tested, because the coefficients of a multinomial choice model are identified only up to a multiplicative scale factor. Strictly, (4) also implies that $\beta_{3i} = \beta_{6i}$, i.e. that the effect of a percentage change in wages is equivalent to that of a percentage change in the tax on wages. We shall not impose this restriction, because (a) we observe taxes on personal income (whose incidence on firms’ wage bills we cannot measure) and (b) our data for wages and for personal income taxes are at different spatial scales. Moreover, for expositional simplicity we shall report results with β_{5i} , β_{6i} , β_{7i} , and β_{8i} each constrained to be equal across sectors, i.e. we assume the effects of taxes on factor inputs and the main effect of N_{ij} (which we shall represent by the empirical variable *sector proximity*) to be the same across sectors. To the extent that they are possible, estimations with sector-level effects of these variables yield qualitatively equivalent results to those reported below. Given the very limited time variation in our data, sector-level identification of β_{7i} , and β_{8i} is not feasible.

¹⁵Silva and Tenreyro (2006) have shown that the Poisson estimator is particularly well suited to log-linear regression specifications that are derived from multiplicative models with potentially heteroskedastic error terms.

distribution. Note that the deterministic part U_{fj} is equal across lots within location j . Then the probability that firm f chooses a lot in location m is given by

$$p_{fm} = \sum_{\ell \in L_m} p_\ell = \sum_{\ell \in L_m} \frac{\exp U_{f\ell}}{\sum_j \sum_{\ell \in L_j} \exp U_{f\ell}} = \frac{L_m \exp U_{fm}}{\sum_j L_j \exp U_{fj}} = \frac{\exp(U_{f\ell_m} + \ln L_m)}{\sum_j \exp(U_{fj} + \ln L_j)}.$$

This is equivalent to a conditional logit model *at the level of locations* with profit function $\pi_{fj} = U_{fj} + \ln L_j + \varepsilon_{fj}$, where ε_{fj} is independent across f and j and follows an extreme-value type 1 distribution. The number of lots L_j in location j can be computed as economically usable land area S_j divided by lot area \bar{s} . The profit function at the level of locations then becomes

$$\pi_{fj} = \tilde{U}_{fj} + \ln S_j + \varepsilon_{fj}, \quad (7)$$

where the constant lot size only affects the intercept of the deterministic part $\tilde{U}_{fj} = U_{fj} - \ln \bar{s}$.

We therefore include usable land area as a control variable in all our estimated regressions. Taken literally, expression (7) suggests that the coefficient on log area should be equal to one. **This restriction is tied to the given fixed variance of the assumed extreme-value type 1 distribution (unlike the slope coefficients in (1) and (5), which are only defined up to a constant scale factor).** We shall test the robustness of our results to the formal imposition of a unit coefficient on log area in the specific model.

Another scaling issue concerns firms, our basic unit of analysis. We adopt two approaches throughout our estimations, by defining the dependent variable alternatively as counts of new firms and as counts of full-time equivalent jobs in new firms. While the former definition follows strictly from the theory (which does not feature different-sized firms), the latter definition takes account of the possibility that larger new firms are more mobile and more sensitive to differences in locational characteristics.¹⁶

3.3.2 Unobserved location-specific effects

Firms' choices may in part be driven by location-specific variables that are unobserved by the econometrician, such as the bureaucratic costs of registering new firms or availability of

¹⁶See Baldwin and Okubo (2008) for a model of location choice by heterogeneous firms in which larger firms are more sensitive to corporate tax differentials than smaller firms.

specialized labor. Omission of these variables can lead to biased parameter estimates on the included regressors. Furthermore, if such unobserved location-specific factors are spatially autocorrelated, the ‘independence of irrelevant alternatives’ (IIA) assumption underlying the conditional logit approach is violated. We therefore also estimate specifications including location-specific fixed effects that control for all unobserved location-specific characteristics. Our baseline model (1) thus becomes:

$$\pi_{ij} = \alpha_3 T_j A_i + \beta' \mathbf{x}_{ij} + \gamma' \mathbf{d}_i + \delta' \mathbf{d}_j, \quad (8)$$

where \mathbf{d}_j is a set of location dummies. This approach no longer allows identification of coefficients on purely location-specific characteristics such as T_j . Since we are interested in this effect as well as in coefficients on certain purely location-specific variables, we estimate our models both with and without location-specific fixed effects, taking the former as robustness tests of the latter.

3.3.3 Identification

In the aggregate, local corporate tax rates and sector agglomeration patterns are both cause and consequence of firms’ location choices. The local stock of firms influences local tax rates through the local tax base or through the political process of local tax setting, and sector-level agglomeration indices are by construction the result of existing firms’ location choices.

Our strategy for avoiding potential simultaneity bias is to study location choices of *new* firms *in narrow sectors* as a function of *lagged* tax rates and agglomeration measures.

While it would seem plausible that existing firms in a local jurisdiction together have influence on prevailing tax rates, we consider it highly unlikely that entrants in a particular narrow sector, location and period exert significant and systematic influence on pre-existing local tax rates. It is important to note in this context that the local jurisdictions of our data are legally bound to set identical statutory taxes across all sectors (see Section 4.1). This allows us to treat tax rates as exogenous not only from the viewpoint of an individual firm but also from that of a cohort of new firms in a particular sector, location and period.

Reverse causality from location choices to agglomeration measures is ruled out by construction, since our agglomeration indices are computed over pre-existing stocks of firms. Furthermore, given the narrow definition of sectors we shall work with, we also feel confi-

dent in abstracting from the possibility that the intensity of spatial concentration could be influenced by the level and spatial distribution of corporate tax burdens.¹⁷

3.3.4 Inference

Some features of our data challenge the assumptions of the standard Poisson model. While the maximum likelihood estimator remains consistent despite these features, we need to take them into account explicitly for correct inference.

We consider three particular issues. First, the Poisson model implies that the expected count, λ_{ij} , is equal to the variance of n_{ij} . This is typically a strong assumption in empirical applications, as the variance often exceeds the expected count (overdispersion), and as economic data often feature a larger number of zero observations than that implied by the Poisson distribution.¹⁸ Second, our models include several explanatory variables that are either purely location-period specific (such as the corporate tax rate) or purely sector-period specific (such as the agglomeration index), while the unit of observation is location-sector-period specific. First, such aggregate variables can lead to standard errors which are seriously biased downward if not correctly adjusted for (Moulton, 1986). Second, even including fixed effects may not account for all the consequences on covariance patterns (Bertrand, Duflo and Mullainathan, 2004). For example, despite our inclusion of sector-period fixed effects, there might be unexplained spatial correlation within sectors. Third, we observe firm startups in two time periods. We cannot exploit this panel structure by including location-sector fixed effects, as the changes over time in our main explanatory are too small for meaningful estimation. However, the likely presence of location-sector random effects needs to be taken into account when estimating standard errors.

We address all three concerns by clustering standard errors in the three dimensions of our data: sector-period, location-period, and sector-location, applying multi-way clustering as proposed by Cameron, Gelbach and Miller (2007). Clustering by sector-period and location-period takes care of the second issue discussed above, clustering by sector-location addresses

¹⁷For models of endogenous agglomeration, driven in part by taxation patterns, see e.g. Ottaviano and van Ypersele (2005), and Haufler and Wooton (2007).

¹⁸In our data, the share of zero observations ranges from 73.3 percent at the two-digit level of sectoral aggregation to 89.7 percent at the four-digit level. In an earlier version of this paper, we also reported zero-inflated panel Poisson (ZIP) estimates (Brühlhart, Jametti and Schmidheiny, 2007). We found the resulting parameter estimates to be very similar. In light of this, and since the ZIP estimator breaks the formal link between the theoretical profit function and the empirical model, we now concentrate on the Poisson estimator.

the third issue, and any of the clusters automatically accommodates the first issue. As we shall show by comparing different variance estimators, multi-way clustering considerably increases standard errors. It therefore represents a conservative approach to inference.¹⁹

4 The empirical setting

4.1 Local taxation in Switzerland

We base our estimations on data for Switzerland. For a number of reasons, the Swiss fiscal system provides an ideal laboratory in which to examine our research question.

Swiss sub-federal jurisdictions enjoy almost complete autonomy in the determination of their tax rates, and, as a consequence, we observe large variations in tax burdens even within the small area covered by Switzerland.²⁰ The Swiss federation consists of three government layers (federal, cantonal and municipal), with each jurisdictional level collecting a roughly similar share of total tax revenue.

In the Swiss fiscal system, corporate taxation is mainly the remit of sub-federal jurisdictions. Cantons and municipalities collect around 65 percent of the total tax revenue raised on corporate income and capital, the remaining 35 percent being raised by the federal government. Hence, differences in corporate taxes across cantons and municipalities matter, and they are large. Figure 1 illustrates this point for consolidated cantonal-plus-municipal corporate income taxes on profits of an average-sized firm with a 2 percent return on capital: the highest tax rate, at 14.1 percent, is almost seven times higher than the lowest rate, at 2.1 percent.

Another convenient element of the Swiss system is that corporate taxation is based on legally binding statutory rates that depend solely on firms' profitability and capital base.²¹

The definitions of these tax bases have been harmonized countrywide by a federal law that

¹⁹In practice, multi-way clustering with large covariance matrices can in some cases lead to negative variance estimates and therefore yield no standard errors (see Cameron et al., 2007, p. 9). We encountered this problem in less than ten percent of regression runs. In these cases, we report instead the most conservative two-way clustered standard errors. In regressions that include location-period fixed effects, we systematically apply two-way clustering (by sector-period and sector-location), as three-way clustering turned out to be computationally infeasible.

²⁰Switzerland has an area of 41,285 square kilometres and a population of 7.5 million. It therefore covers about twice the area, and hosts about the same population, as the US state of Massachusetts.

²¹A number of cantons and municipalities also levy taxes on real estate and/or on property transactions. These taxes too are not specific to certain firms or sectors, and they are of limited importance. Country-wide data on such taxes paid by corporations are not available. We can, however draw on statistics released by the canton of Basel Stadt, according to which, over the period 2001-2004 revenue from property taxes raised from corporations was equivalent to a mere 4 percent of the revenue raised from corporate income and capital taxes.

has been in force since 1993 and that foresees no firm-specific or sector-specific regimes except for some clauses to avoid double taxation of holding companies.²²

The sole exception to equal treatment across firms and sectors at the sub-federal level is that some (mainly industrial) firms can be offered tax rebates for a maximum of ten years after setting up a new operation. No systematic data are made available for cantonal and municipal exemptions granted, but available evidence suggests that they affect less than 4 percent of new firms.²³ Furthermore, as tax holidays at the federal level are contingent on exemptions granted at the cantonal level, cantons and municipalities have a strong incentive to grant exemptions more generously if concurrent tax exemptions are granted by the federal government. Since the eligibility for federal tax holidays is restricted to certain legally defined ‘lagging’ regions, we can partly control for the incidence of local tax exemptions by including an indicator variable for those municipalities that belong to eligible regions.

Sector-specific taxation exists at the federal level (for value-added tax, excise taxes and import duties), but all cantonal and municipal taxes imply identical treatment across sectors. In addition, we work with narrow sector definitions: the average share of a sector in terms of both employment and firm counts across our sample municipality is 0.48 percent with the three-digit sector definition, and 0.21 with the four-digit definition. Municipalities dominated by one or a small number of sectors are exceedingly rare. Cases in which a sector accounts for more than 10 percent of municipal *employment* represent 0.39 percent of all observations at the three-digit level, and 0.25 percent at the four-digit level. Cases in which a sector accounts for more than 10 percent of the municipal *firm count* represent 0.26 percent of all observations at the three-digit level, and 0.02 percent at the four-digit level.

²²The official title of the law is “Bundesgesetz über die Harmonisierung der direkten Steuern der Kantone und Gemeinden”, adopted by the federal parliament on 14 December 1990. Special tax treatment applies to farming, but we omit agricultural activities in our estimations. For firms with operations in several cantons, the exemption principle holds. Double-taxation agreements define the allocation of profits using formula apportionment, mostly based on wage bills, capital or sales (see Feld and Kirchgässner, 2003, p. 135, for an illustrative example). Note that the exemption principle combined with formula apportionment provides an incentive for firms to respond to tax differentials via physical location choices rather than through creative accounting.

²³According to published government replies to parliamentary questions in the cantons of St. Gallen and Lucerne, 59 and 35 temporary tax exemption agreements were granted respectively by these two cantons over our sample period 1999-2002. Relative to the number of firms created in those cantons and years, this represents 3.8 and 3.6 percent respectively. These percentages must be considered upper bounds, as some exemptions are granted to existing firms that undertake significant restructuring projects.

4.2 Data sources

We draw on data from three main sources. First, the Swiss Federal Statistical Office has collected information on every newly created firm annually since 1999.²⁴ The main use of this data set is as the source of new firm counts per municipality and economic sector (n_{ij}), one of our dependent variables. The data also report full-time equivalent employment numbers in these firms, which we shall exploit for the construction of our alternative dependent variable, sector-municipality counts of jobs in new firms. We use data for the years 1999-2002. The database also offers information on the municipality in which the new firm is located and on the firm's main sector of activity in terms of the three- and four-digit sectors of the European NACE classification system.²⁵ The data set records as new firms all market-oriented business entities that have been founded in the year concerned and are operating for at least 20 hours per week. New entities created by mergers, takeovers or breakups are not counted, nor are new establishments by existing firms. A foreign firm's first Swiss branch, however, counts as a new firm. Observed firm start-ups undoubtedly represent a mixture of births through resident entrepreneurs best modeled by the latent-startup approach and of births by non-resident (Swiss or foreign) investors best modeled by the footloose-startup approach.

Our second source of data is the multiannual census of all firms located in Switzerland, also carried out by the Federal Statistical Office.²⁶ The census records establishments (of which there can be several per firm) and attributes them to a NACE sector according to their self-declared principal activity. We use data for the surveys of 1998 and 2001, containing information on location, sector of activity and employment, to construct our agglomeration variables.

Finally, we have assembled a municipality-level dataset on local taxes and other control variables from a variety of sources.²⁷ We use these data for our measures of corporate and personal income tax burdens, factor prices, public expenditure, and proximity to markets. The data were collected for 1998 and 2001, covering the 213 largest municipalities. The mean population of our sample municipalities is 17,367, for a mean total area of 20.2 square

²⁴The statistical office's title for this project is "Unternehmensdemografie" (UDEMOMO).

²⁵We retain only sectors that pertain to the private sector. Furthermore, sectors for which no firm births are observed in either year are dropped from the dataset. This leaves us with 133 sectors at the three-digit level and 242 sectors at the four-digit level.

²⁶The statistical office's title for this project is "Betriebszählung" (BZ).

²⁷For a detailed description of the data on municipal taxes and other municipal attributes, see Brühlhart and Jametti (2006).

kilometers.²⁸

4.3 Variables used

4.3.1 Dependent variable

We run all of our regressions for two waves of firm creations. Counts of new firms set up over the period 1999-2000 (or of full-time equivalent jobs created by those firms) are assigned to control variables for 1998, and those created over the period 2001-2002 are assigned to control variables for 2001. The average number of new firms (n_{ij}) per location and three-digit industry is 0.486 and 0.440 in our two sample periods respectively. The nationwide sample totals are 13,768 new firms in 1999-2000 and 12,465 new firms in 2001-2002. These firms created 27,756 full-time equivalent jobs in 1999-2000 and 22,454 in 2001-2002. See Table 1 for summary statistics of the dependent and independent variables.

4.3.2 Explanatory variables: baseline model

Our main explanatory variables in the baseline model (2) are local corporate taxes (T_j), sectoral agglomeration economies (A_i) and, most importantly, the interaction of those two effects. We represent local corporate taxes via a *tax index*, a revenue-weighted average of consolidated municipal and cantonal profit and capital taxes. The index is calculated separately for 1998 and for 2001. Corporate income tax schedules are progressive in most municipalities. Hence, we collected statutory corporate income tax rates for three representative levels of profitability, 2, 9 and 32 percent, and took the mean of these three rates as an index for the corporate income tax. As capital taxes are generally proportional, we collected statutory capital tax rates for a firm with the median capital base. To compute the *tax index*, we normalized the profit-tax index and the capital tax rates by subtracting the mean and dividing by the standard deviation for each of the two sample years, and we weighted them by the respective importance in terms of tax revenue. Hence, the *tax index* has mean zero by construction. Note that none of our qualitative results hinges on the exact construction of the tax index.²⁹

²⁸Due to the small size of our sample jurisdictions, we feel confident in abstracting from within-jurisdictional heterogeneity. Duranton, Gobillon and Overman (2007) provide a careful treatment of this issue based on data for English Local Authorities (which, on average, cover areas that are 18 times larger than our Swiss sample municipalities).

²⁹Results for our baseline model using statutory tax rates instead of the *tax index* can be provided on request. Data on the distribution of profitability rates, capital bases and revenue shares of different tax instruments are taken from publicly available federal tax statistics.

Agglomeration economies are not directly observable. In steady state, however, sectors subject to strong agglomeration economies will be more spatially concentrated than sectors subject to weak agglomeration economies (or to net dispersion economies). Hence, we compute spatial concentration indices using the definition proposed by Ellison and Glaeser (1997), which controls for differences in firm numbers across sectors in quantifying the extent of geographic clustering. We refer to this variable as the *EG index*. Our coefficient of interest in the baseline model (2) is α_3 , the effect of the interaction between the *tax index* and the *EG index*. We compute this interaction by multiplying the two indices after mean-differencing the *EG index* (the *tax index* has mean zero by construction). Thereby, the interaction term has a mean of zero, which allows us to interpret the estimated coefficient on the *tax index* as the effect of taxes for a sector with average spatial concentration.

A number of control variables are included (\mathbf{x}_{ij}). In order to allow for cost factors affecting firm profits, we control for the prices of labour and of real estate.³⁰ *Wage* reports average monthly wages per sector and region in the year 2000, while *property price* stands for the average selling price of a representative family home by municipality in the year 2002.³¹ In the baseline model, we interact both these price variables with the *EG index*, as we may expect equivalent effects of agglomeration economies for the importance of factor prices to those we hypothesize for local tax rates: the stronger are sector-specific agglomeration economies, the less sensitive firms' location decisions should be, other things equal, to differentials in factor prices across municipalities.

As personal income taxes may affect firms' location choices in addition to corporate taxes, we also include the variable *income tax*, which represents the canton-averaged statutory cantonal-plus-municipal personal income tax rate for a median-income representative household. We choose this measure, which is invariant across municipalities within each canton, because distances within cantons are sufficiently small to allow easy commuting among municipalities. Hence, income taxes in the particular municipalities where firms are located

³⁰We assume that the price of capital is equalized across Swiss municipalities.

³¹*Wage* is available from the Swiss national statistical office at a level of sectoral aggregation corresponding roughly to one-digit NACE, and at the level of regions comprising several cantons ("Grossregionen"). It is thus assumed that relative wages are constant over our sample period, among subsectors and within regions. *Property price* is available from the consultancy firm Wüest & Partner. Since commercial property prices are not collected at a sufficient level of detail for our purpose, we employ prices of private property as the best approximation. It is assumed that relative wages and property prices did not vary significantly over our sample period.

would not be the relevant measure.³² Similarly, we control for *public expenditure*, computed as canton-averaged municipal-plus-cantonal expenditures on the main spending items from the viewpoint of private-sector firms: education, public safety and transport. Again, selecting only municipality-specific expenditure would not represent the relevant variable, as Swiss municipalities are sufficiently small for a large share of commuting to take place between rather than within municipalities.³³

Proximity to same-sector firms is the key cost factor stressed in the economic geography literature - either as a source of technological spillovers, specialized production factors and intermediate inputs, or as factor affecting the intensity of local competition. We therefore compute the variable *sector proximity* for each sector, municipality and base year. This variable counts the number of existing same-sector firms in a municipality plus the corresponding number firms in all other Swiss municipalities, weighted by the square of Euclidian distance between the geographic municipality centers.³⁴ In regressions featuring employment as the dependent variable, *sector proximity* is computed over sector-municipal employment rather than over firm counts. The main demand-side control variable is *market potential*, which, for each municipality, is defined analogously to *sector proximity* as the inversely distance weighted average income across all Swiss municipalities.³⁵ As a simple complementary measure, we also include *distance to highway*, the road distance to the nearest access point to the highway network. This variable, unlike *sector proximity* and *market potential*, has the advantage of measuring accessibility without implying that the relevant economic space ends at the national border. Summary statistics on these variables are provided in Table 1. We furthermore include a dummy variable for *assisted municipalities*, which are defined as lying within a region identified by federal law as eligible for temporary tax exemptions for newly created firms. Finally, we control for *area*, for consistent estimation given unequally sized locations (see Section 3.3.1). In defining this variable, we consider only built-up and constructible

³²We have also performed our estimations by replacing the canton-averaged tax rate on a median-income household by (i) the municipal median-income tax rate, (ii) the canton-mean maximum (i.e. high-income) tax rate, and (iii) inversely distance weighted averages of municipal tax rates. All our qualitative results are unaffected by these variations in the definition of *income tax*.

³³Using a broader measure of public expenditure does not affect any of our qualitative results but leads to somewhat weaker effects of the expenditure variable.

³⁴Square weights are chosen in view of the existing evidence on spatial decay functions based on intranational commuting patterns (e.g. Harsmann and Quigley, 1998) We approximate intra-municipal distances by $2/3\sqrt{\text{area}/\pi}$.

³⁵Municipal incomes are estimates reported by the Swiss federal statistical office for 1992, the latest available year.

surfaces.

4.3.3 Explanatory variables: specific model

Estimation of our specific model (5) requires a subset of the variables used in the baseline model, measured in logs and, in the case of tax rates, $\ln(1 - t_j)$. An important difference is that we need to identify the relevant tax rates. We therefore consider statutory tax rates on representative tax payers for taxes on corporate income (t_j^π), capital (t_j^k) and personal income (t_j^w).³⁶ Specifically, we retain two corporate income tax rates: *median-profit corporate tax*, which applies for a firm with 9 percent profitability, and *high-profit corporate tax*, which applies for a firm with 32 percent profitability. Swiss fiscal statistics suggests that median profitability in our sample period was around 9 percent, whereas 32 percent represents profitability of a firm at the lower end of the highest sextile. We consider both these tax rates alternatively, as it is *a priori* similarly plausible that firms consider the tax burden on a “representative” case or that they mainly pay attention to the high end of the tax schedule.

5 Results

5.1 Some preliminary Illustrations

Before reporting econometric estimates, we provide some illustrative examples of our central result. Figures 2, 4 and 6 show maps of the geographic distribution of establishments in three sectors: software development and consulting (NACE 722), production of clothing (NACE 182), and brokerage and wealth management (NACE 671). These sectors are shown in increasing order of their measured geographic concentration, the first two sectors having below-average *EG indices*, and the third sector displaying an *EG index* above the overall average of 0.016. The corresponding Figures 3, 5 and 7 plot counts of new firms, scaled by area, in 1999-2000 over the 1998 *tax index* of all municipalities that have witnessed firm births over the relevant period. We observe that the relationship between taxes and firm births becomes gradually less negative as we move down the list of sectors. While the relationship between taxes and firm births is negative and statistically significant at the one-percent level for soft-

³⁶Based on fiscal statistics, we take a representative capital stock as 176,000 and 181,000 Swiss francs respectively in 1998 and 2001, and we consider canton-averaged income tax rates on a household with two children and a taxable annual income of 73,000 and 75,000 Swiss francs respectively in 1998 and 2001. One Swiss franc traded for 0.63 US dollars on average over our sample period.

ware development and consulting, this relationship turns positive, although not statistically significant, for brokerage and wealth management. These examples illustrate our main point: the more spatially concentrated a sector, the less firm births in that sector are deterred by high local corporate taxes (or attracted by low taxes).

5.2 Baseline Model

We begin by estimating our baseline model, which includes fixed effects for sector-periods but not for municipalities. This specification allows us to identify coefficients on municipality-specific variables.

As a preliminary exercise, Table 2 reports estimates of specification (2) without any controls (\mathbf{x}_{ij}) except for the scaling variable *area*. The model is estimated for both three- and four-digit sector definitions, and for firm and employment counts as alternative dependent variables. We report four different standard errors: unadjusted Poisson, bootstrapped, panel-robust and three-way clustered.³⁷ The table shows clearly that the choice of inference method matters: our preferred three-way clustered standard errors are up to 24 times larger than the unadjusted Poisson standard errors and up to five times larger than the robust standard errors. We focus on these most conservative standard errors in the remainder of our analysis.

Our key parameter of interest, on the interaction of the *tax index* and the *EG index* (α_3), is positive even in the pared-down regressions shown in Table 2, thus confirming our central hypothesis. The effect is larger both in terms of the coefficient size and of statistical significance when we take employment counts as the dependent variable. This suggests that tax differentials and agglomeration matter more for location choices of large firms than for those of small firms.

Table 2 reports positive main effects of the *tax index*, which is clearly not plausible and calls for a fuller specification that controls for determinants of firm location other than corporate taxes and agglomeration. This is what we do in Table 3, where we estimate the full baseline model (2). The table also reports estimates of the baseline model with inclusion of municipality-period fixed effects. These regressions serve as a robustness check for unobserved municipality-period-specific features that might affect location choices and thus bias the base-

³⁷The bootstrap draws sector-period groups and thus ignores clustering along the location-period and the location-sector dimension. Panel-robust standard errors are computed taking sector-period groups as relevant units of observations. They are thus equivalent to one-way clustering by sector-period.

line estimates. When including municipality-period fixed effects, we can no longer identify the main effect of the *tax index* nor of any other sector invariant regressors.

The coefficient on the *tax index* now turns negative, as expected. The interaction term with the *EG index*, however, remains positive. This configuration is robust across all eight specifications reported in Table 3. We have furthermore examined the robustness of this central result (a) to considering statutory corporate tax rates instead of the *tax index*, (b) to including interactions with sector-level shares of existing firms located in urban regions, and (c) to controlling for stocks of existing firms by municipality and sector. Our qualitative findings turned out to be robust to these alternative specifications as well.³⁸

The estimated coefficients on the included control variables largely conform with expectations. We find consistently negative coefficients on *wage*, *income tax* and *distance to highway*, and positive coefficients on *public expenditure*, *area*, *sector proximity* and *market potential*.³⁹ Interactions of *wage* and *property price* with the *EG index*, reflecting the prediction that firms' sensitivity to factor prices other than taxes is also lessened as they experience stronger co-location economies, return positive estimated coefficients, in line with expectations.

The only unexpected result is the estimated coefficient on *property price*, which is statistically significantly positive. The most plausible explanation for this result is that *property price* correlates with unobserved location-specific features that are attractive to new firms and to some extent capitalized in property prices. We interpret this result as suggesting omitted variables at the municipality level, which supports our consideration of municipality-period fixed effects as a robustness check on the full baseline model (columns 3, 4, 7 and 8 of Table 3). The estimated interaction coefficients in these regression runs are somewhat smaller in magnitude, but they are again consistently positive; and statistical significance is found in all four regression runs. In line with expectations, significantly positive effects are also found on the interaction of property price with the *EG index*, and on *sector proximity*. Neither the main effect of wage nor its interaction with the *EG index*, however, is statistically significant. This is not very surprising, as we observe wages only at the level of large regions and not of

³⁸Results tables can be provided on request.

³⁹We have also experimented with variables capturing municipality-level human capital (measured by average educational attainment). Educational attainment turns out to be strongly collinear with *market potential*, and inclusion of human-capital variables can yield negative estimated coefficients on *market potential*. This is consistent with models in which centrality (generally implying location in an urban area) is attractive mainly because of the access it offers to a skilled workforce and in spite of local congestion diseconomies (see, e.g., Combes, Duranton and Gobillon, 2008). None of the remaining coefficients are significantly affected when human capital is included.

individual municipalities.

Our baseline estimations therefore confirm the hypothesis we seek to test: location choices of firms in more spatially concentrated sectors are less sensitive to tax differentials.

The main effect of taxes is statistically insignificant in all of the regressions shown in Table 3. This need not, however, imply that corporate taxes are never a statistically significant deterrent to firm births - it only means that corporate taxes do not have a statistically significant impact on location choices of firms in sectors with average agglomeration intensity.⁴⁰ We illustrate this in Figure 8, where we report the total coefficient on the *tax index*, $\hat{\alpha}_2 + \hat{\alpha}_3 A_i$, based on our estimates in the first column of Table 3, across values of the *EG index* (A_i). Confidence intervals are computed via the delta method from the three-way clustered standard errors shown in Table 3. Our estimates imply a negative effect of corporate taxes on new firm counts and new firm employment for the all values of the agglomeration index up to one standard deviation above the mean. This range covers over 90 percent of sample sectors. Figure 8 also shows that the negative effect of taxes is statistically significant for sectors with an *EG index* of up to half a standard deviation below the mean. Approximately a quarter of sample sectors fall within this range and are thus found to be statistically significantly deterred by corporate taxes.

5.3 Specific model

Table 4 reports our estimates of the specific model (5), with all coefficients constrained to be equal across sectors (allowing us to report them in the table).⁴¹

Again, we find that high corporate taxes deter firm births. (Note that the positive coefficient estimated on $\ln(1-t_j^\pi)$ implies that the effect of the tax rate is negative.) However, statistical significance on the corporate tax coefficient is found only for the specifications featuring the tax rate on high-profit firms. It therefore appears that location choices are primarily guided by the top end of the corporate tax schedule.

The effects of *income tax* and *capital tax* are also estimated to be negative. Capital taxes appear to be considerably less important than corporate and personal income taxes. This is probably explained by the fact that capital taxes play a relatively minor role in the Swiss fiscal

⁴⁰Recall that, in these baseline regressions, we use standardized agglomeration measures, such that our sample sector with mean agglomeration has a standardized *EG index* of zero.

⁴¹For the specific model, we report only regressions with firm counts as the dependent variable. Corresponding estimations for employment yield qualitatively equivalent results.

system, accounting for a mere three percent of consolidated tax revenues at the sub-federal level (whereas corporate income taxes represent some twelve percent and personal income taxes about two thirds of total sub-federal revenues).

Conversely, agglomeration effects, measured here as the coefficient on the interaction between *sector proximity* and the *EG index*, are positive throughout and statistically significant in the two three-digit estimation runs. The remaining controls perform in line with expectations: a large *area*, high *sector proximity* and high *market potential* raise the number of new firms, while a high average *wage* appears to be detrimental. The estimated coefficients on *property price* are not statistically significant, which is again suggestive of a dual role played by this variable, both as a factor price (which deters firm births) and as a positive but imperfect correlate of unmeasured locational attractions (which promote firm births), thus supporting inclusion of municipality fixed effects to test the sensitivity of our parameters of main interest. This is what we do in the final two specifications shown in Table 4. These estimates confirm the importance of agglomeration effects: firms tend to locate in the proximity of other same-sector firms, and this tendency is stronger in more strongly agglomerated sectors.

In Table 5, we examine the robustness of the regressions presented in Table 4 to two specification changes that take the empirical model closer to the theory. The first panel of Table 5 reports estimates of a specification that allows for sector-specific coefficients on *wage*, *property price* and *market potential*, and thereby gets closer to expression (5).⁴² Once more, our main results stand: high corporate and personal income taxes depress firm births, whereas firms in highly agglomerated sectors, measured by the *EG index*, choose locations with high *sector proximity*. Our qualitative results also hold once we force the coefficients on *area* to unity, as suggested by the empirical model in Section 3.3.1. These estimates are reported in the second panel of Table 5.

5.4 Quantitative interpretation

A comparison of our estimates for the baseline and specific models is not trivial, due to three fundamental differences of specification. First, the deterministic part U_{ij} in the profit function

⁴²We report results for interactions of these three variables with dummies for one-digit sectors. Due to space constraints, we report test statistics for the joint significance of each set of coefficients rather than listing all the individual estimates. Interactions with dummies for more disaggregated sectors do not substantially alter our results.

is assumed to be additive in the baseline model (U_{ij}^B in equation 1), while it is multiplicative in the specific model (U_{ij}^S in equation 5). Second, the baseline model uses a *tax index* while the specific model uses statutory corporate tax rates. Third, the interaction effect between taxes and agglomeration is explicit in the baseline model but only implicit in the specific model.

An economically meaningful interpretation of the parameters that also allows a comparison across the two models is nonetheless possible, based on the conditional-logit approach. Unlike linear regression coefficients, conditional-logit parameter estimates cannot be interpreted as marginal effects. However, the parameters serve to predict the probability that a firm from sector i chooses location j : $P_{ij} = \frac{e^{U_{ij}}}{\sum_k e^{U_{ik}}}$. Marginal effects are obtained by differentiating this expression. Thus, the impact of a marginal increase in taxes in location j on the probability that a sector- i firm picks location j is given by:

$$\frac{\partial P_{ij}}{\partial T_j} = (\alpha_1 + \alpha_3 A_i) \cdot P_{ij}(1 - P_{ij}) \quad (9)$$

in the baseline model, and by

$$\frac{\partial P_{ij}}{\partial t_j^\pi} = -\frac{\beta_2}{1 - t_j^\pi} \cdot P_{ij}(1 - P_{ij}) \quad (10)$$

in the specific model.

These marginal effects depend on all estimated parameters and all variables, through their dependence on P_{ij} . They therefore differ across locations, sectors and periods. To compute comparable marginal effects, we average all variables that vary across sectors except for the *EG index*, in order to isolate the interdependence of the tax effect with agglomeration economies from other cross-sectoral differences. We then visualize the marginal effects for a representative municipality (Montreux), focusing on the first wave of data (1998-2001).⁴³

Figure 9 shows the marginal tax effect as a function of the agglomeration index for both the baseline and the specific model. In order to make the magnitude of the effect comparable, the effect is drawn for a one standard deviation increase of the respective tax variable.

The first panel of Figure 9 is very similar to Figure 8, the difference being the scale of the vertical axis, which now has a clear economic meaning. Again, we see that the tax effect

⁴³Montreux is a representative municipality in the sense that it is close to the sample average both in terms of market access and of population size.

is strongest and statistically significant in sectors with small agglomeration economies. In more agglomerated sectors, the impact of taxes shrinks and finally even turns (insignificantly) positive.

Take a concrete example. The software sector (NACE 722) is relatively dispersed, with a standardized *EG index* of -0.46 corresponding to the twentieth percentile of our sample distribution. Our estimated marginal tax effect for the software sector is -0.0004 and significantly different from zero. This means that the probability that a new firm locates in Montreux increases by 0.04 percentage points if Montreux were to lower its *tax index* by one standard deviation. The seemingly small effect needs to be compared to the average choice probability (P_j), which is 0.3 percent for that municipality. Hence, a one-standard-deviation reduction in the *tax index* will raise the predicted number of new software firms in this municipality by $0.04/0.3 = 13.3$ percent. Conversely, the agricultural machinery sector (NACE 293) is relatively agglomerated, with a standardized *EG index* of 0.09 corresponding to the eightieth percentile of the sample distribution. Our estimated marginal tax effect for the agricultural machinery sector is -0.00026, one third smaller than for the software sector and not significantly different from zero. Hence, while a one-standard-deviation reduction of the tax index will attract 13.3 percent additional software firms, it attract only 8.7 percent additional firms belonging to the agricultural machinery sector. Expressed differently, moving from a sector in the top agglomeration-intensity quintile to a sector in the bottom agglomeration-intensity quintile increases the tax sensitivity of firms' location choices by 50 percent.

The second panel of Figure 9 shows the corresponding marginal tax effects for the specific model. Although the tax coefficient does not directly depend on the *EG index* here, the marginal effect of taxes still does. Intuitively, when agglomeration economies get stronger, proximity to existing same-sector firms becomes more important to firms' location choices, and hence local tax burdens become relatively less important. The second panel of Figure 9 illustrates how the marginal effect of taxes is stronger for more agglomerated than for less agglomerated sectors - again in line with our basic hypothesis. Two differences, however, also become apparent. On the one hand, the marginal effect of taxes is now statistically significantly negative for all values of the *EG index*. This strengthens the conclusion that, other things equal, corporate taxes act as a deterrent to firm location. On the other hand, however, the relationship between the tax effect and the intensity of agglomeration economies

emerges as very weak in this specification. In the representative example illustrated in Figure 9, the marginal tax effect falls (in absolute terms) from -0.000259 in the software sector to -0.000247 in the agricultural machinery sector. Hence, moving from a sector in the top agglomeration-intensity quintile to a sector in the bottom agglomeration-intensity quintile increases the tax sensitivity of firms' location choices by some 5 percent only. Given the heavy constraint imposed on this effect by the specific model, which allows it to feature only implicitly, this probably can be considered a lower-bound estimate of the dampening impact of agglomeration economies on the tax sensitivity of firms' location choices.

6 Conclusions

Drawing on a firm-level dataset for Switzerland and employing fixed-effects count-data estimation techniques, we find that firm births on average react negatively to corporate tax burdens, but that the deterrent effect of taxes is weaker in sectors that are more spatially concentrated. This finding supports the validity of recent theoretical results suggesting that agglomeration economies can reduce the importance of tax differentials for firms' location choices and thereby lessen the intensity of corporate tax competition even if technological and administrative barriers to capital mobility are low.

In a sense, this research constitutes but the first step in a full evaluation of the prediction that agglomeration forces mitigate 'race-to-the-bottom' tax competition. Although tax competition is often at its fiercest when targeted at new firms, it could be useful to explore how tax differentials affect not just births but the entire life cycle of firms, including expansions, contractions and deaths. In future work it will furthermore be interesting to study whether policy makers recognize the differential impact of fiscal inducements across sectors and effectively seek to tax agglomeration rents, and whether this effect is strong enough to have a noticeable impact on the evolution of statutory corporate tax burdens.⁴⁴ Finally, it would be interesting to distinguish between, on the one hand, spatial concentrations due to exogenously given endowments and, on the other hand, agglomeration of essentially footloose firms attracted to each other by various types of externalities. In theory, the latter type of agglomeration forces can, depending on parameter values, intensify tax competition rather

⁴⁴Two recent papers, by Charlot and Paty (2007) and Solé-Ollé and Jofre-Monseny (2007), suggest that local taxes are indeed positively correlated with local agglomeration economies.

than mitigating it.

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